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Forever gender equal and child friendly? Intrahousehold allocations to health in Finland before the Nordic welfare state

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ABSTRACT: The literature on intrahousehold allocation in European history has typically built on bargaining models originating from Amartya Sen and the South Asian “missing girls” paradigm, testing hypotheses of male earner bias. Often, a 50/50 benchmark has been used, assuming any skew in spending meant discrimination. This study combines external measures of variation in morbidity by age, sex and season with analysis of household health expenditure in Finland in the 1920s. The results suggest that money largely followed sickness rather than gender or earnings. This supports an emerging literature challenging bargaining models and suggesting that significant historical differences may have existed between world regions.

1. Introduction

Research on intrahousehold allocation has identified discrimination against female children across a range of goods, including health services. Typically, this has been explained by expected returns from future income or the bargaining power of current earners (in development economics, e.g. Qian 2008; Burgess & Zhuang 2003; Sen 1984, pp. 374-376; in economic history, e.g. Horrell & Oxley 2013; Horrell, Meredith & Oxley 2009; Moehling 2005). Historical discussion has been most lively on industrializing Britain. In several articles over the years, Sara Horrell, Deborah Oxley et al. (2013; 2009; 1999) have argued male earner bias and the applicability of “Senian” bargaining models, using anthropometric as well as household expenditure data. Bernard Harris (2008) and Andrew Hinde (2015) have questioned the interpretation specifically of female excess mortality in certain age groups as evidence of discrimination, suggesting that an analysis of causes of death and regional patterns rather supports epidemiological explanations. Katherine Lynch (2011) has identified factors working against neglect and infanticide of girls in Europe as a whole, ranging from the existence of wage labor opportunities to religious norms, arguing that the continent was inherently different. McNay, Humphries and Klasen (2005) have conducted a regional analysis within Britain, and found a complicated landscape where both epidemiological and socioeconomic conditions mattered, but space for regional differences based on local culture and norms also remained. Indeed, the findings of intrahousehold literature are not homogenous even in South Asia, and more equitable, even pro-girl, allocation patterns have also been detected in specific cases (e.g. Luke & Munshi 2007; on education, Kingdon and Theopold 2008; Munshi and Rosenzweig 2006; Himaz 2010).

The issue is important as a growing literature suggests that long-run differences in economic and social development between regions may partly have had deep roots in microinstitutions related to gender, family and demography (e.g. De Moor & van Zanden 2009; Diebolt & Perrin 2013; Alesina, Giuliano & Nunn 2013). This paper looks at a Nordic country before the Nordic social model, and the case of intrahousehold allocations to health. The results indicate that health expenditure was apparently to a significant degree guided by variation in objective health risks rather than intrafamiliar power relations or gender discrimination. Furthermore, under tighter financial constraints at the bottom of the income distribution, children were given preference over adults. This is not self-evident, and contradicts previous literature emphasizing male earner bias that often deprived women and children. The implication is that the famous modern Nordic gender equality and child friendliness could have had long roots.

In the 1920s after gaining independence from Russia, Finland was a distinctly poor and peripheral country, with a GDP per capita of about 66-75 % of Sweden's and only 50-60 % of the UK's in 1929 (Prados de la Escosura 2000) as well as around 70 % agrarian workforce and an urbanization rate below 20 % (Koskinen et al. (eds.) 2007, p. 323; Kaukiainen 1981, p. 40). Epidemiological transition was still ongoing, with contagious disease, particularly tuberculosis, accounting for high share of mortality among children and youth (Koskinen & Martelin 2007, pp. 181-182). In 1940 a government committee estimated that a fifth of the population was still quantitatively malnourished (Komiteamietintö 1940:5). Interwar social policy was still characteristically 'unscandinavian' and residual in nature, with no mandatory health insurance in place before the 1960s (Kettunen 2001; Mattila 2011). At the same time, GDP per capita more than doubled from 1920 to 1940 (Hjerppe 1988, pp. 215-216), and a modest expansion of modern services was creating new kinds of employment opportunities for women. Less wealthy socioeconomic groups were more likely to send girls than boys to secondary education when they could, because the immediate opportunity cost in lost earnings was lower and relative expected returns higher. However, enrolment was still generally low (Saaritsa & Kaihovaara 2016).

A recurring methodological problem in intrahousehold allocation literature across disciplines is lack of attention to demonstrable variation in needs – and therefore the welfare returns to allocations – by age and gender. Nutritional requirements for boys and girls may differ at different ages, particularly when activities differ (Floud, Fogel, Harris & Hong 2011, pp. 41-49). Different returns to education on gendered labor markets may explain seemingly unequal schooling decisions by families (Saaritsa & Kaihovaara 2016; Himaz 2010; Goldin 1998, pp. 362-363). In the case of medical expenditure, there is clear evidence that risk of illness, and hence the need for care, varied significantly by age and sex in historical populations. This makes the 50/50 expenditure used by some researchers (e.g. Simonsson 2004; Burgess & Zhuang 2003, p. 9) an unsatisfactory benchmark for equity. To gain a better understanding of the logic of allocations, proxies for age and sex specific morbidity need to be combined with an analysis of household expenditure.

This article uses household budget data from Finland in the 1920s to analyze the responses of differently positioned households to illness. The data makes it possible to measure unpaid as well as paid consumption, and map variation according to income, social group and type of habitat. Detailed information on household composition makes it possible to draw on other demographic data and medical literature to determine the presence of groups at higher than average risk in terms of health, and to apply an Engel model used in development economics to discern patterns of intrahousehold allocation. Quarterly records make it possible to exploit seasonal variation in morbidity, the regional prevalence of certain types of disease, and to apply panel regressions. The findings show a clear connection between metrics of objective risk and household allocations as well as consistent prioritization towards highest risks, and ultimately children, connected with tighter financial constraints.

Section 2 describes relevant features of health services in early 20th century Finland. Section 3 presents the core data, a household budget survey from 1928, and discusses the incidence of services – in a sense, the "tacit policy" (Rainwater, Rein & Schwartz 1986, pp. 22-23) emerging from the system – by social group and type of habitat. Section 4 constructs proxies for sickness, for which there is no direct data, and looks econometrically at the characteristics of responses to elevated morbidity risk by households in different sorts of positions. Section 5 concludes.

2. The patchwork of care

Most medical services available in Finland in the 1920s had to be paid for at some rate, but there was significant variation. The makeup of Finnish health institutions was an outcome of layers of history, creating data and differences in access which can be exploited in the analysis. Since Swedish rule in the 18th century, a sparse network of state-employed district doctors (*piirilääkärit*) spanned the country. In the first attempt ever to provide public medical services (Saltman & Dubois 2004, p. 24), their core mandate was public health and combatting epidemics. There were at most 57 district doctors in the beginning of the 20th century, serving areas with varying size and population, and regularly reporting on health indicators. They did see patients as well, and according to a code dating from the 1830s were obliged to help the "poor" according to "their oath and conscience" (Mattila 2006, pp. 18-20, p. 23; Harjula 2012, p. 34.).

The most important public provider were municipal doctors (*kunnanlääkärit*). Hired from the 1880s onwards, they offered treatment to locals, but had to be paid according to ability. Differences were great. In the capital Helsinki, the city hired several doctors to low-income neighborhoods with a mandate to treat everybody without screening. Other cities had special "doctors of the poor" on payroll. Elsewhere, free consultation required a statement from poor relief officials, deterring patients with humiliating board hearings and a debt to the municipality (Harjula 2012, pp. 34-36; Hannikainen 2001). Unlike competing claims on municipal resources from new compulsory schooling and poor relief legislation in the beginning of the 1920s, the provision of health services was not required by law. The majority of the rural, poor municipalities, with local politics dominated by farmers unwilling to pay more taxes, simply did not hire one. In 1932, there were 254 municipal doctors and 601 municipalities in Finland (Lavonius 1959, p. 103; ”Kaupunkien ja kuntien lukumäärät 1917-2013”; Kuntaliitto.).

Another notable source of subsidized supply were fringe benefits from employers. The nascent wood processing industry created a largely rural and isolated workforce (Alapuro 1985, p. 64; Tilastollinen päätoimisto 1930, p. 105), and had to provide essential services independently. According to a 1927 survey, 28 % of c. 200 factories hired doctors and 35 % had a contract with an outside party, typically a municipal doctor (Kuusi 1931, pp. 940-941). Rural factory workers were, in fact, privileged in terms of health care, able to frequent a doctor more often than an average citizen (Jaakkola 1996; Harjula 2012, p. 34.). Also white collar employees could enjoy similar benefits.

[TABLE 1 HERE]

The share of predominantly urban private practitioners increased during the 1920s, exceeding that of municipal doctors. It has been estimated that in interwar Helsinki, earnings had increased enough to make this an alternative for workers as well (Siipi 1962, pp. 348-349). Table 1 provides a breakdown of health professionals by sector. Meanwhile, in the late 1930s, a fifth of rural municipalities were still completely without a doctor of any sort, public or private (Harjula 2012, p. 33) and two thirds had no hospital services (Mattila 2011, p. 95). A significant part of the rural poor lived without a realistic chance to see a medical professional, left to traditional healers, enlightened clergy and quacks (e.g. Hako & Tuomi (ed.) 2000; Saarivirta, Consoli & Dhont 2010, p. 29; Mattila 2006, pp. 18-19; Piela 2006.).

3. The data

The main source is the primary material of a 1928 government Cost-of-Living Study (CLS), targeting the non-agrarian population. It contains 954 household budgets from 14 cities and 15 smaller population centers, mainly rural industrial communities. One in four survey households were from Helsinki (pop. 227 000 in

1928), followed by the South-Eastern and South-Western coastal hubs of Viipuri (8 % of CLS households, pop. 54 120) and Turku (7 %, pop. 63 918) (SVT XXXII: 14, p. 4; Tilastollinen päätoimisto 1930, p. 11). Most of the remaining 15 localities were towns and villages dominated by a single or a few industrial employers.

Probabilistic sampling had not yet become internationally recommended best practice (Desrosières 1991; Kruskal & Mostelles 1980). In larger cities, the selection of households was delegated to local authorities. The study was to focus on "workers proper", adding lower and higher officialdom and families "belonging to the so-called middle class". It was instructed that "typical occupations" of each locality should be covered (SVT XXXII: 14, pp. 3-4). The data had an "intact" family bias, consisting almost solely of households with two parents and children. The internationally favored diary method used (e.g., Halbwachs 1913, pp. 29-32) was based on households keeping accounts on their income and expenses themselves for a period of one year, monitored by a "bookkeeping advisor". This excluded those without will or capacity to do so. The result was deemed to represent "the cost of living of a certain kind of an elite" (SVT XXXII: 14, p. 3).

However, for Helsinki, the average family size conformed to that in the city, while parents of underage children tended to be older than average adults, and fewer were elderly (Saaritsa 2008b, pp. 318-320). The worker household heads of the CLS had median incomes comparable to the adult male population of the city according to municipal taxation. The data also contained workers with low and insecure income, including a small number of past or present recipients of poor relief, a sign of major hardship in the 1920s. Several families suffered income shocks during the year, some due to illness (Saaritsa 2011, pp. 105-127). The sample is thus not limited to a 'labor aristocracy'. The selection issues are standard for similar datasets in economic history (e.g. Horrell & Oxley 2000, pp. 38-39; Horrell & Oxley 1999, pp. 497-498; Moehling 2001, pp. 932-933; Emery 2010, p. 76). For this analysis, the ability to represent the whole population is less relevant than the ability to discern families with different ability to consume medical services.

Out of the three groups defined by the investigators, workers (*työläiset*) included people working for wages with or without vocational training. White-collar employees (*toimenhaltijat)* consisted of low-ranking government and municipal officials like tram drivers and policemen, and private sector employees like shop assistants and skilled mechanics who ”on the basis of their income and standard of living” were considered similar. Two thirds were employed by the public sector, while the figure for workers was only 11 % -- still an overrepresentation.[1] The elite (*virkamiehet*[2]*)* included high-ranking government officials, professionals and ”comparable” cases, slightly over half from the public sector (SVT XXXII: 14, pp. 9-11).

The quarterly summary cards in the archives of Statistics Finland (Tilastokeskus, Tilastoarkisto, K09e Kulutustutkimukset) include household composition, earnings and a variety of expenditures. Health expenditure in cash was recorded as a sum spent on fees and medication. A useful feature of the accounting framework was the category of “expenditure in kind”. It contained a cash estimate of the value of goods and services that had been received in kind rather than purchased with money – services from a municipal doctor or a factory hospital, for instance. The monetary value was supposedly based on local market rate, and should be considered with caution in light of the limited methodological repertoire of the time. The figures still enable gauging who received unpaid services, who did not, and what the rank order of different kinds of households was. Most importantly, the data allows comparing households in terms of the consumption of health care across social groups, income levels and localities.

Table 2 reveals expected and unexpected differences between the social groups. Worker families had more children and were marginally less dependent on the earnings of the household “head”. A quarter of workers had more than one child below five years of age, while for others, the share was approximately 15 %.

[TABLE 2 HERE]

Unsurprisingly, wealthier groups used more money on medical services. More interestingly, the incidence of unpaid services does not seem to have been progressive; white-collar employees actually consumed slightly more unpaid care than workers, albeit the difference may not be statistically significant. The estimated amount of unpaid services consumed was lowest among the elite, which likely resorted to private practitioners.

Table 3 breaks down the consumption of unpaid services by type of locality and social group. Although cell size gets small, it appears that the unpaid consumption of the urban white collar employees was higher in relative terms than that of urban workers. As they were unlikely to obtain more free services from municipal doctors, potential sources included employee benefits or sickness funds that offered services directly. The employees of the State Railways, altogether 28 % of the white collar workers surveyed, were a key group. Their share out of those who consumed unpaid services was over half outside the cities (11 / 19), and as high as three quarters in the cities (24 /32). This overrepresented group, easy to reach and recruit for the study, obviously enjoyed popular medical benefits.

[TABLE 3 HERE]

As for workers, differences between rural and urban localities underline the importance of the benefits offered by the outlying industrial colonies, obliged to maintain their human resources independently. Over a third of the rural workers had consumed free services during the year, and around 20 % had exceeded 100 Finnish *markka*s (FIM), which was about 6 % of an average monthly worker salary. In contrast, only a small percentage of urban workers had received any unpaid services at all. This would suggest that the municipal services, on which they mainly had to rely for free care, in fact had a rather limited role in average consumption.

4. Time to see a doctor?

How did households with varying ability to consume medical services respond to sickness? The usefulness of available treatment as such has been a contentious issue in the history of health (McKeown 1979; Szreter 1988). Concurring with findings like those by Winegarden & Murray (2004; 1998), early 20th century medical view still maintained that doctors had an important role in the treatment of common illnesses. In absence of antibiotics, early diagnosis was important, as a condition could quickly progress to serious complications. Many poxes ideally required hospitalization. Mortality from diphtheria decreased radically since the adoption of serum therapy invented in the 1890s, which could also be used to temper measles and scarlet fever. Pre-antibiotic treatment of infections required care: e.g. diphtheria could cause strangulation, a condition needing a surgical operation to keep the respiratory tract open (Söderström 1933).

There were thus good reasons to assume that seeing a doctor when sick had positive health effects, and it was rational to seek treatment. While some therapies now known to be ineffective or even harmful were practiced – and certainly had been practiced as part of traditional folk medicine for centuries (Strandberg 2012) – for an argument on discrimination in the *responses* of households, the ultimate optimality or effectiveness of the measures is not decisive.

4.1 Risk groups

The source contains no direct data on illness apart from some scattered mentions. However, it is possible to proxy variation in morbidity through indirect indicators. Firstly, it is possible to utilize the presence of groups with elevated morbidity risk in households, using information on household composition by age and sex.

While differences in health expenditure by sex have been used as such to argue discrimination within households, there is robust evidence of variation in morbidity between sexes at different ages. In reference populations considered relatively equal, males tend to have higher mortality than females across age groups, and specifically the morbidity of young male children is usually higher than that of female children (cf. Klasen 1998, pp. 434-437). For example a large cohort study on Finnish children born in 1987 indicates 22 % higher mortality and consistently higher morbidity among boys aged 0-7 compared to girls of the same age (Gissler, Järvelin, Louhiala & Hemminki 1999). Although negligence and infanticide have at times skewed this, higher mortality among young boys has been a common finding also across populations in preindustrial Europe and Asia, and there is research supporting biological explanations to early male frailty (Alter et al. 2009, p. 329). In Finland in 1928, the mortality of boys aged 0-4 was approximately 16 % higher than that of girls, and larger than in any age group below 60 years of age (SVT VI Väestötilastoa 75.). Young male children were thus plausibly a group with higher needs than females in terms of medical expenditure.

On the other hand, female excess mortality from late childhood till the end of the reproductive years is an established historical demographic pattern, which has disappeared over time in advanced countries (e.g.Alter et al. 2009, pp. 327-328). When found coinciding with industrialization and modernization processes, explanations have been sought in changing patterns of livelihoods, bargaining positions, as well as working and living conditions. While the most important proximate cause has almost universally been identified as pulmonary tuberculosis (TB), it has been suggested that its progress was in turn triggered by gender specific changes, such as growing intrahousehold discrimination in nutrition (for an overview, see Lynch 2011, pp. 258-260). This reasoning has been contradicted by empirical findings showing similar excess mortality in highly diverse circumstances, including rural, non-industrial and preindustrial Europe and Asia (Alter et al. 2009; Henry 1989, p. 196). In the UK, regions where intrahousehold discrimination has been otherwise identified do not match the regions of female excess mortality (Hinde 2010). An alternative line of inquiry works on the hypothesis that young females were more susceptible to TB for physiological reasons (Janssens, Messelink & Need 2010, p. 92; Henry 1989, p. 196; Ehrenreich & English 1976). The phenomenon itself is in any case robust, and its historical demise could be linked with general decline in disease mortality as well as progress in gender equity (Harris 2008).

In Finland, female excess mortality, measured by sex ratio in risk of dying obtained from life tables, appeared clearly in ages 11 through 15 in the 1880s[3], peaked in the first decade of the 20th century, remained evident in the 1920s, and was in decline by the 1930s (appendix figure 1). As usual, its rise and fall took place in the context of steadily declining mortality for *both* sexes, but with different paces (cf. Harris 2008). Data from life tables on relative risk of dying by age over the 1920s yields a profile that supports the hypothesis of higher risk for young boys and a switch around ages 11-17 to the detriment of girls (figure 1). The measurement does not reflect the exact year 1928, which may be significant as epidemiological conditions were constantly changing. Population statistics report deaths by age and sex for 1928, but the size of the at-risk population for calculating mortality is not available. An estimate based on the raw ratio of deaths by sex is likely to be indicative, and can be adjusted with the small difference between raw ratio and mortality in the census year 1930. The result comes surprisingly close, main differences being higher volatility of the male excess in younger age groups, and the onset of the female excess from age 12 rather than age 11. In addition to stochastic variation, these could involve a degree of measurement error.

[FIGURE 1 HERE]

Cause of death by disease is also available for pre-defined age groups for 1928, and the role of TB in driving the female excess seems evident (appendix figure 2). Notably, the simple removal of TB deaths would remove the excess, although such a static exercise disregards the redistribution among competing causes of death that would occur (Hinde 2010).

Regardless of the verdict on the role of discrimination, girls in these age brackets can be considered a higher-risk group than boys. However, it is notable that the *absolute* levels of risk, included in figure 1, were dramatically lower for both sexes in this age group than among young children. This matters when risk is proxied with presence of people, as the implied variation in needs occurred at a much lower frequency.

We can also assume that the morbidity of ageing individuals would have been higher than that of prime age workers. While family bias in selection led to an underrepresentation of older age groups, slightly over 10 % of the households still had members aged 55 and over (see Saaritsa 2008b, p. 319). Within the age group, the presence of older individuals was skewed towards women (79 as opposed to 53 men). Among the over 60s and the over 65s, there were roughly three times more women than men. While the onset of elevated morbidity and mortality occurred earlier for males, aging itself was to a significant extent a gender specific risk for women in a relatively pensionless society.

It is difficult to determine whether any variation in intrahousehold allocation was proportional to the variation in health risks, or whether elements of discrimination remained. Particularly the relationship between different types of health problems correlated with mortality and levels of health expenditure was likely to be heterogeneous, making it hard to assess the adequacy of response. In any case, while some studies have mechanistically suggested that allocations of over 50 % to one sex are automatically evidence of discrimination (Simonsson 2004; Burgess & Zhuang 2003, p. 9), in addition to cultural factors or bargaining power, also variation in morbidity should be considered as an explanation. In the context of differing needs, unequal allocations could in fact be consistent with equalizing returns to well-being for all family members (cf. Klasen 1998, pp. 434-435, fn 11).

The effect of risk groups on the health expenditure of differently positioned households can be analyzed with a Working-Leser Engel model, widely applied in development economics (Deaton 2000, pp. 231-233; Burgess & Zhuang 2003; Himaz 2010). It operates through modelling linear Engel curves for the consumption of goods and adding household demographics on the age, sex and number of household members to identify gendered differences in allocation patterns.

In the standard model:

$$w_i = \alpha_i + \beta_i \ln (x/n) + \eta_i \ln n + \sum_{j=1}^{J-1} \gamma_{ij} (n_j/n) + \delta_i z - \mu_i$$

w_i is the share of commodity i – in this case, health – of total household expenditure, x is the total household expenditure, n is the household size, n_j is the number of household members in the age-sex group j, while z stands for the vector of relevant household characteristics and μ_i is the error term. The impact of the sex of children is tested through comparing γ_{ij} for relevant female and male age groups. The key indicator for gender differences in intrahousehold allocation is the F-test on the difference between boys and girls in the same age group (Deaton 2000, p. 234). Regardless of the significance of the coefficients, the order of the point estimates indicates the direction of this difference.

Two models are run by social group, an Engel specification operating with shares and a specification using raw amounts of people and cash. The causes for the discrepancies in results are elaborated further with nonparametric methods.

Table 4 shows significant gender differences in allocation only in the group of workers, which was on average under tighter constraints than the higher socioeconomic groups. In the age group 0-10 the difference in the coefficients between the sexes suggests allocations in favor of boys. The results of the F-tests hold across models, while there is no indication of male bias in other age groups. This is consistent with the notion that allocations responded to risks.

[TABLE 4 HERE]

It would, however, be important to show that the same applied to females when health odds were against them. A mirroring result is indeed suggested by the findings on older girls, but the results are somewhat tenuous. A modest conventional threshold of significance is only achieved with the model based on amounts for the 11-17 age group. Furthermore, sensitivity analysis reveals that the results are not robust to incremental changes in the boundaries of the age categories.[4]

This type of data is inevitably noisy. A nonparametric exploration (figure 2) maps the relationships of the relevant variables for worker households two-dimensionally using a locally weighted regression (see DiMatteo 1998), superimposing two options for the age brackets. Comparing left side with right side, the figures would seem to neatly match the results on the OLS models in the sense that the gender difference in the 11-17 group is clearer when amounts are used instead of shares. This would suggest that the smaller difference in the Engel models is driven by something in the *denominators* of the Engel variables, which depend on *total* household composition. The graphs imply that some factor, possibly related to characteristics of the other members of households of certain size, interferes with the explanatory variable intended to capture the presence of children of certain age and sex, making the results less concerning.

[FIGURE 2 HERE]

On the other hand, the graphs for the 12-17 grouping, based on the 1928 death data included in the figure, show that there is a narrowing of the gender gap compared to the 11-17 groupings, mirroring loss of significance in the OLS. Descriptively, the pattern of dominance of female over male effect, which matches health risks, still remains. The order of the graphs is the same with the other possible age groupings discussed, with LOWESS lines on females always higher than those on males. Finally, the absolute levels of risk were far lower in the older than in the younger age group, making the expected link between people and health expenditure weaker. It can therefore be concluded that there was observable traction towards girls in expenditure at risky ages, although not robustly identifiable with a linear Engel specification.

In the over 55s age group, there is a significant difference in favor of females in the Engel specification. In the model using amounts the p-value is .109, which can be considered suggestive. A pro-female estimate would concur with the higher death risk of the more aged females in this group. With values extracted from life tables, the average combined worker-household-specific death risk for males over 55 was 2,25 per 1000 and that for females 3,54 per 1000 – a nearly 60 % difference. While not unequivocal, the implied response of health expenditure to higher female risk in this group contradicts male bias.

Since indicators of health risk are available, is it possible to control for differences in needs between groups directly, and discern whether any potential discrimination remained? Life tables provide a specific risk of dying for each age and sex in the 1920s. Health expenditure is still measured at household level, so the need for an Engel-style specification remains. A feasible application is attaching a risk of dying to each individual in a household, and summing these risks by age-sex group. These group risks can then be related to the total sum of risks in the household in order to measure the weight of groups in the overall risk burden of a household, resembling the procedure carried out previously with household demographics.

However, the death risk based indicators turn out to be highly correlated with those based on numbers of people. This is particularly the case in the 11-17 age group (r>0.96 for both sexes and models), while the connection is attenuated in the 0-10 age group.[5] This evidently reflects the lower absolute variation in risk in the 11-17 bracket. In terms of F-tests and significances, the regressions consequently yield substantively identical results (see appendix table 1). The difference in the 0-10 age group is no longer significant when amounts of expenditure are regressed on death risk per 1000, while in the over 55s age group a very strong pro-female difference would be apparent with both specifications.

Should the persistence of significant differences be taken as evidence of the existence of discrimination after objective needs are controlled? There are good reasons to abstain from straightforward interpretations. Firstly, as pointed out, the gender order of the coefficients varies between age groups, and would seemingly indicate discrimination shifting from females to males by age (cf. Saaritsa & Kaihovaara 2016, pp. 79-80, 82-83). Secondly, risk of dying and demand for medical services cannot be assumed to have had a homogeneous relationship across sex and age categories. For instance, a large share of the early excess mortality of male children was caused by deaths from birth defects during the first month of life. From about age three onwards, deaths by accident and violence started to explain a significant part of the male excess. Each of these categories of causes of death could have led to a specific pattern of medical spending, possibly generating less demand than diseases if there was less nonfatal morbidity related to them. However, also pneumonia, a common complication of childhood diseases, was more elevated among males, as were unspecified "other diseases". The female excess in the 11-17 bracket, then again, was largely driven by TB (SVT VI Väestötilastoa 75, table 27). This could have caused lumpy expenses to wealthier families that were able to send members to private sanatoriums, while an insufficient range of largely ineffective costless facilities were available for the very poor, with not much on offer in between (Harjula 2015, pp. 109-112, 205).

The argument made is therefore essentially ordinal. While it is plausible that groups with higher risk of dying had on average higher needs in terms of medical expenditure, the specific level of need cannot be controlled with available indicators precisely enough in order to identify and measure ”pure” discrimination. Discrimination cannot be excluded, although the shifts in preferred gender over age would complicate the story. However, the findings do support the conclusion that objective risks were significant in triggering allocations. Certainly, there is no evidence of consistent male bias in this data.

In light of the group differences in the ability to consume paid and unpaid care, it also seems that an analysis based on high-risk groups works systematically less well for households that were either more able to spend, like the elite, or more effectively covered by fringe benefits, like the white collars. All statistically significant patterns identified pertain to workers only. This could be considered somewhat surprising – after all, one could have expected that families with better resources would have been able to respond more observably to their members' needs. Yet similar findings are generated using other morbidity proxies, and can be given a rational interpretation.

4.2 Seasonality

Sickness varied over time within the year as well, and on this, indicators measuring morbidity instead of mortality are available. District doctors reported observed cases of contagious disease in their area by month. Apart from clearly seasonal afflictions like influenza, many diseases exhibited less regular cycles. Measles epidemics, for instance, did not take place every year in the entire population, although annual outbreaks might have occurred in the cities (cf. Waris 1934, pp. 99-107). This is why same-year data is useful. Figure 3 presents the total number of cases and breakdown for most important diseases during 1928.

[FIGURE 3 HERE]

The graphs indicate typical seasonal variation, with an elevated disease burden in the winter months (cf. Floud, Fogel, Harris & Hong 2011, p. 181.). The pattern was driven by influenza. From January to April-May, however, measles and whooping cough were clearly more prevalent than during the rest of the year, January and February being the hardest months. Towards the end of the year, there was also a minor increase in scarlet fever, which at this time could lead to weeks of isolation in a hospital and up to a month and a half's absence from school (Söderström 1933, p. 113).

Figure 4 presents the quarterly disease burden and health expenditure in cash by CLS households. The expenditures of the elite and the workers, both of whom relied predominantly on paid services, varied in concert with morbidity, with the more low-income workers spending at a lower range. An exception to the pattern were white collar employees, likely because of the documented high frequency of access to unpaid care.

[FIGURE 4 HERE]

Quarterly data enables the use of fixed effects regressions to estimate the effect of seasonal variation on medical expenditures in differently positioned households. The standard time and entity fixed effects estimator operates through regressing, for each individual household, the deviations from annual mean of health expenditure on similarly time-demeaned quarterly expenditure. Quarterly dummies are then added to capture the effect of different seasons. In the equation:

$$(y_{it-}\bar{y}_i) = \beta(x_{it} - \bar{x}_i) + \eta Q1 + \gamma Q2 + \delta Q4 + (\epsilon_{it} - \bar{\epsilon}_i)$$

y_{it} is the health expenditure for household i on quarter t, $\bar{y}_i$ is the annual mean of the health expenditure for household i over all quarters, x_{it} is the log household expenditure for household i on quarter t, $\bar{x}_i$ is the annual mean of the log household expenditure for household i, Q1 to Q4 are dummies for the respective quarters with the least diseased quarter 3 left as reference, and ϵ stands for the error term similarly transformed. Time-invariant factors, such as household demographics, cannot be included, but are controlled through the fixed effects (e.g. Wooldridge 2003, pp. 461-463). The dependent variable applied is health expenditure rather than consumption share in order to avoid distortion by seasonal variation in the total consumption denominator, but the results are substantively unaffected by this. As data on in-kind consumption is only available at the annual level, the dependent variable only contains paid consumption.

Table 5 compares the effect of seasonal variation on health expenditure in worker households divided into four splines on the basis of household income from poorest to richest, and in elite households. The elite is a particularly interesting reference, because the group met its needs predominantly through paid, private services, but did not suffer from tight budget constraints. The results show that workers with low income only increased their expenditure significantly on the first, most difficult quarter. Moving up the income distribution, the number of significant quarters grows and the significance of the estimates increases. The first quarter is consistently identified, but more affluent workers spent significantly more on all the three quarters that were more disease-stricken than the summer months. Also the elite spent significantly more on two out of the three more burdened quarters. For the relatively poorest group, then again, only the first quarter had a significant effect.

[TABLE 5 HERE]

The emerging group differences can be read as an indication of *prioritization* within households. All groups across the distributions of income and coverage were able to consume some medical services, but when and for whom was a different matter. In households belonging to wealthier social strata or with good access to unpaid services, demographic risk groups were not statistically discernible. This suggests expenditure did not have to be prioritized as clearly, and could be spread more evenly. In terms of seasonality, the poorest households only responded to the period of highest morbidity. In the context of the seasonal labor market of the late 1920s, this was *not* the period after the summer when the savings of workers were highest, so the timing cannot be explained simply by mechanical spending on emerging needs until money ran out (Saaritsa 2011, pp. 102-106). Under tight constraints, highest needs were apparently rationally prioritized. With increased resources, expenditure spread more evenly across seasons of elevated morbidity to respond to more diverse needs, and the spending pattern of workers started to converge with that of the elite.

4.3 Children first

In terms of *who* was prioritized, many of the results hint that children were likely to be at the top of the list. In the risk group analysis, the most consistent finding was on young boys. Differences in the expenditure on boys and girls may or may not have been commensurate with differences in morbidity. The tentative indications of opposite flows in age groups where girls had higher risk still agree with a pattern of rational response to the health needs of all children regardless of sex when finances were constrained. Seasonally, the robustly significant first quarter was the one with clearly the highest incidence of childhood diseases (measles, whooping cough). However, this aggregate pattern can be insufficient evidence of targeting children. Literature has identified near-identical seasonal patterns in morbidity, with the beginning of the year consistently as the peak, using historical data far removed from Finland in 1928, which suggests that this feature is more universal (Harris et al. 2012, p. 726, p. 728; Floud, Fogel, Harris & Hong 2011, p. 181).

It is possible to test the hypothesis more credibly by exploiting regional heterogeneity and available data on the prevalence of specific diseases. Published statistics enable distinguishing childhood diseases from conditions affecting only or also other age groups by drawing on contemporary medical literature (Söderström 1933). While some uncertainty is inevitable, the two most significant illnesses in this regard, measles and influenza, were straightforward. The previous affected almost exclusively young children and gave survivors lifelong immunity, while the latter was everybody's seasonal plague. Both displayed high prevalence and strong variation during 1928. Morbidity can furthermore be measured at the level of mainland Finland's eight provinces (*lääni*).[6] On this level, epidemic patterns varied more, and the peaks for children's and adults' diseases did not always occur simultaneously. Influenza and adult diseases peaked in the second quarter in several provinces, while childhood diseases more consistently peaked at the beginning of the year, and particularly in the case of measles remained down for the rest of the year. Even if provinces were quite large and not optimal units from an epidemiological viewpoint, linking provincial morbidity with household budgets yields results.

Table 6 presents the relationship of the regional prevalence of different types of disease to health expenditure by income and social group within the year using a fixed effects estimator. Quarterly regional morbidity is measured by cases reported in the province per 1000 inhabitants in 1928, while seasonal time dummies are omitted for collinearity. Estimates are presented for two models: one with all contagious disease classified into childhood diseases[7] and diseases affecting only, or to a significant degree, also adults[8]; and one with measles and influenza only.

[TABLE 6 HERE]

The results reveal a pattern resembling the findings on seasonality. At the bottom of the income distribution, they imply a strong and highly significant relationship with childhood illnesses only. Moving higher, the statistical effects become smaller and weaker, and in the top income quartiles among the workers, adult conditions take over. The estimates are even clearer when focusing only on measles and influenza. In both models, they are similar between the market-reliant but financially unconstrained elite, and the group of workers as a whole. While there is no substantive interpretation for the disappearing effect of childhood illnesses among wealthier workers, the pattern of enhanced focus on children towards the bottom of the distribution is evident.

5. Conclusions

The findings of this article support the conclusion that under tight financial constraints, a fundamental determinant of health expenditure within early 20th century Finnish households was need rather than power or gender. Econometrics of household expenditure shows consistent behavioral patterns along the distributions of income and entitlement. Further away from higher incomes and well protected groups, only indicators of high probability of sickness– young boys prone to illness, the exceptional disease burden of the first months of the year – attained robust significance, suggesting rational and altruistic prioritization within households. In terms of gender, there are a number of indications that differences in spending on boys and on girls in certain ages were caused by differences in the likelihood of being ill. In the age groups where the mortality of girls became excessive over that of boys in the 1920s, nonparametric graphs suggest that spending followed, and some, if not all, standard regressions validate this. Analysis of responses to different types of disease strongly suggests that particularly children were being prioritized at the cost of adults. Childhood diseases yielded the clearest and most consistent responses among low-income, market-dependent groups, while variation in adult-related sickness caused a response only among slightly more affluent households.

The prioritization of children of different sexes according to medical needs within households is not a trivial finding. Previous research on intrahousehold allocation has typically highlighted preference for workers and earners, explained either by bargaining position or by collective necessity to maintain work capacity. In 19th century England, this evidently led to widespread discrimination of women and children in terms of nutrition (Harris 2008, pp. 45-46; Horrell, Meredith & Oxley 2009; Horrell & Oxley 2013). However, it appears that in interwar Finland, sick children in worker families did get to see a doctor even if it cost money, also when there was not much to spend from.

The results are relevant for the evolving discussion of female excess mortality, lately focused on the UK. While explanations based on discrimination have remained prominent, e.g. the work by Harris (2008) and Hinde (2010) has been qualifying this approach. The Finnish findings are at odds with the discrimination hypothesis in at least two ways. Firstly, if money for treatment followed morbidity also in the age group where female excess mortality was discernible, it would be hard to reconcile this with the notion that the health risk itself was caused by neglect within the same households that were allocating their scarce resources to medical care. Secondly, recent analysis has shown that girls from worker families in this age group were, on average, given more resources for education than boys (Saaritsa & Kaihovaara 2016), contradicting notions of parental exploitation or disinvestment.

The inclusion of historically determined variation in health needs by age and sex into the empirical analysis seems essential in order to make inferences about discrimination within households. While more elaboration is required, it is also necessary to take seriously the possibility that there were economic, social, cultural and other factors that made, for instance, the British, Nordic, German or South Asian gender and family systems fundamentally different from each other in this respect in given historical times (cf. Lynch 2011; Klasen 1998, p. 436). Deep, persistent patterns may also have carried over to later social development, in this case resonating with high public investment into children and relative gender equity during Finland's postwar era as a Nordic welfare state. Indeed, the development of a modern public health care system in Finland famously began with children and postnatal clinics, with adult health lagging badly behind up to the 1960s (Harjula 2012).

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Tables and figures

Table 1. Doctors by sector and mobile nurses in Finland, 1920-1928.

	Doctors, municipalities			Doctors, state			Doctors, private		Doctors, total
	Munic. doctors	Hospi-tals	Other	Dist-rict	Hospi-tals	Other	Hired	Private practice	
1920	171 (26 %)	38 (6 %)	21 (3 %)	57 (9 %)	78 (12 %)	101 (15 %)	29 (4 %)	167 (25 %)	662 (100%)
1924	208 (27 %)	52 (7 %)	17 (2 %)	54 (7 %)	81 (11 %)	111 (14 %)	33 (4 %)	212 (28 %)	768 (100%)
1928	244 (26 %)	62 (7 %)	24 (3 %)	57 (6 %)	112 (12 %)	104 (11 %)	53 (6 %)	270 (29 %)	926 (100%)

Sources: SVT XI Lääkintölaitos 37, taulu 1; SVT XI Lääkintölaitos 41, taulu 1; SVT XI Lääkintölaitos 45, taulu 1.

Table 2. CLS 1928. Descriptive statistics and health expenditure by social group.

	Workers (N=581)		White collar (N=242)		Elite (N=131)	
	Mean (s.d.)	% of expen-diture (s.d.)	Mean (s.d.)	% of expen-diture (s.d.)	Mean (s.d.)	% of expen-diture (s.d.)
Family size	4.93 (1.84)	-	4.62 (1.72)	-	4.33 (1.22)	-
Children aged 0-4	0.85 (0.95)	-	0.64 (0.81)	-	0.71 (0.78)	-
Household income	28983.19 (8920.61)	-	37099.48 (12854.03)	-	58369.07 (20548.11)	-
Main income	20531.56 (6373.19)	76.91 (16.59)	26256.80 (6414.67)	80.21 (17.59)	39047.87 (13050.68)	78.46 (18.67)
Health, total	385.27 (374.80)	1.43 (1.29)	529.34 (673.16)	1.58 (1.95)	1019.45 (1179.38)	1.94 (1.93)
Health, paid	361.95 (369.56)	1.32 (1.20)	503.25 (667.45)	1.49 (1.93)	1009.83 (1173.64)	1.91 (1.92)
Health, unpaid	23.32 (90.54)	0.11 (0.50)	26.09 (76.58)	0.08 (0.20)	9.62 (56.63)	0.024 (0.16)

Source: CLS data.

Table 3. CLS 1928. The consumption of unpaid medical services by social group and type of locality.

	Urban			Rural		
	Workers (N=374)	White collar (N=187)	Elite (N=110)	Workers (N=207)	White collar (N=55)	Elite (N=21)
Received, over 0 mk	20 (5.4 %)	32 (17.7 %)	3 (2.7 %)	74 (35.8 %)	19 (34.6 %)	3 (14.3 %)
Received, over 100 mk	9 (2.4 %)	10 (5.4 %)	1 (0.9 %)	36 (17.4 %)	7 (12.7 %)	2 (4.8 %)

Source: CLS data.

Table 4. Health expenditure and risk groups according to life tables 1921-30. OLS regressions by group.

Model 1: Shares (Engel curves)				Model 2: Amounts			
Dependent: Health exp., % of total	Workers (N=581)	White collar (N=242)	Elite (N=131)	Dependent: Health exp., mk	Workers (N=581)	White collar (N=242)	Elite (N=131)
ln family size	0.09 (0.62)	1.43 (1.54)	3.94 (2.89)	ln family size	1814.91 (800.92)	1036.12 (1545.48)	8701.20 (2394.29)
ln exp per capita	0.30 (0.30)	1.07 (0.72)	1.97 (1.15)	ln exp per capita	504.40 (205.01)	539.98 (452.50)	2400.77 (767.74)
Share of females 0-10	0.85 (0.84)	2.55 (1.30)	3.40 (2.36)	N of females 0-10	-172.26 (77.44)	10.21 (109.45)	-411.10 (248.61)
Share of males 0-10	1.36 (0.81)	3.23 (1.75)	2.14 (2.59)	N of males 0-10	-141.96 (81.14)	28.55 (125.31)	-556.88 (221.32)
Share of females 11-17	0.42 (0.87)	1.10 (1.87)	2.55 (3.17)	N of females 11-17	-173.47 (83.49)	-87.71 (121.51)	-645.75 (290.36)
Share of males 11-17	-0.06 (0.80)	1.92 (2.03)	2.34 (3.30)	N of males 11-17	-211.11 (80.61)	-52.32 (120.48)	-604.70 (424.10)
Share of females 18-29	1.21 (0.86)	2.33 (1.66)	-0.50 (3.52)	N of females 18-29	-139.58 (88.64)	-47.50 (133.93)	-644.49 (358.79)
Share of males 18-29	1.15 (0.70)	0.72 (1.36)	1.56 (1.73)	N of males 18-29	-72.77 (48.79)	-76.76 (113.73)	-405.84 (200.61)
Share of females 30-54	1.59 (0.98)	5.61 (2.43)	1.84 (3.54)	N of females 30-54	-173.92 (116.32)	196.57 (216.45)	-508.03 (296.44)
Share of females over 55	2.96 (1.08)	2.98 (2.89)	3.27 (4.17)	N of females over 55	-78.67 (61.98)	37.91 (176.68)	-887.19 (345.90)
Share of males over 55	-0.51 (1.27)	2.53 (2.70)	-4.45 (2.22)	N of males over 55	-170.14 (73.21)	87.39 (216.61)	-442.70 (218.96)
Constant	-0.28 (1.71)	-5.67 (4.36)	-10.57 (8.21)	Constant	-2917.71 (1346.39)	-2467.08 (2914.58)	-15982.99 (5201.32)
F-test gender diff 0-10	2.90*	0.34	0.30	F-test gender diff 0-10	4.71**	0.69	0.35
F-test gender diff 11-17	2.15	0.23	0.94	F-test gender diff 11-17	4.07*	0.34	0.91
F-test gender diff over 55	4.96**	0.90	0.21	F-test gender diff over 55	2.74	0.81	0.28
R^2	0.07	0.13	0.08	R^2	0.07	0.12	0.10

Source: CLS data. Standard errors clustered by locality in parentheses. ***=p<.01, **=p<.05, *=p<.10 (F-tests only).

Table 5. CLS 1928. Health expenditure and seasonal variation within households. Workers by income quartile and the elite. Fixed effects, OLS regressions by group.

	Workers (N=581)					Elite (N=131)
	All workers	I income quartile (N=146)	II income quartile (N=145)	III income quartile (N=145)	IV income quartile (N=145)	
ln total expenditure	78.72*** (17.55)	80.54*** (25.49)	89.74 (27.65)	64.35 (42.03)	68.00** (30.36)	369,25** (137.87)
Quarter 1	49.22*** (7.61)	45.82*** (13.20)	50.58** (15.06)	27.56** (11.63)	70.65*** (14.79)	147.48*** (34.62)
Quarter 2	29.37 (11.07)**	8.83 (10.22)	6.06 (13.12)	16.34* (8.75)	85.96*** (25.95)	150.54** (57.94)
Quarter 4	9.81 (9.47)	-8.84 (9.72)	4.77 (11.74)	-1.49 (4.91)	45.87** (20.56)	48.97 (49.74)
Constant	-614.37 (154.45)	80.54 (25.49)	-709.31 (237.74)	-482.02 (371.17)	68.00 (30.36)	-3268.65 (1288.75)
R^2	0.05	0.04	0.03	0.03	0.03	0.07

Source: CLS data. Dependent variable: health expenditure, mk. Standard errors clustered by locality in parentheses. ***=p<.01, **=p<.05, *=p<.10.

Table 6. CLS 1928. Household health expenditure and intra-year provincial morbidity patterns. Fixed effects, OLS regressions by group.

	Workers (N=581)					Elite (N=131)
	All workers	I income quartile (N=146)	II income quartile (N=145)	III income quartile (N=145)	IV income quartile (N=145)	
MODEL 1						
ln total expenditure	53.19*** (18.47)	44.03* (21.68)	64.49** (23.22)	43.59 (41.85)	63.14 (38.76)	325.56** (118.80)
Childhood diseases	14.79*** (5.02)	22.37*** (5.71)	17.17** (7.31)	11.47* (6.70)	8.50 (14.52)	52.88* (26.78)
Other diseases	3.65 (3.3)	-7.22 (5.34)	1.18 (3.31)	-0.72 (1.74)	15.55** (7.11)	14.82 (10.55)
Constant	-404.92 (164.31)	-324.32 (182.44)	-507.50 (202.59)	-305.74 (371.57)	-502.70 (356.56)	-2905.83 (1104.89)
R^2	0.05	0.03	0.01	0.01	0.05	0.05
MODEL 2						
ln total expenditure	60.56*** (19.86)	57.16** (21.23)	77.51*** (25.01)	43.31 (45.68)	69.49* (38.34)	334.60** (127.27)
Measles	18.61*** (5.13)	28.66*** (5.84)	25.91*** (8.63)	9.67 (8.43)	12.53 (7.91)	56.03** (22.29)
Influenza	6.21** (2.61)	-3.83 (4.72)	3.66 (2.46)	1.85* (1.07)	16.79*** (4.07)	24.37*** (7.13)
Constant	-461.40 (176.97)	-422.24 (177.29)	-611.54 (217.73)	-297.03 (404.44)	-555.02 (353.24)	-2955.81 (1186.11)
R^2	0.05	0.03	0.01	0.02	0.05	0.06

Notes: Morbidity measured as reported cases per 1000 inhabitants. *Sources:* SVT XI Lääkintölaitos 45; Suomen Tilastollinen Vuosikirja 1930, p. 10; CLS data. Dependent variable: health expenditure, mk. Standard errors clustered by locality in parentheses. ***=p<.01, **=p<.05, *=p<.10.

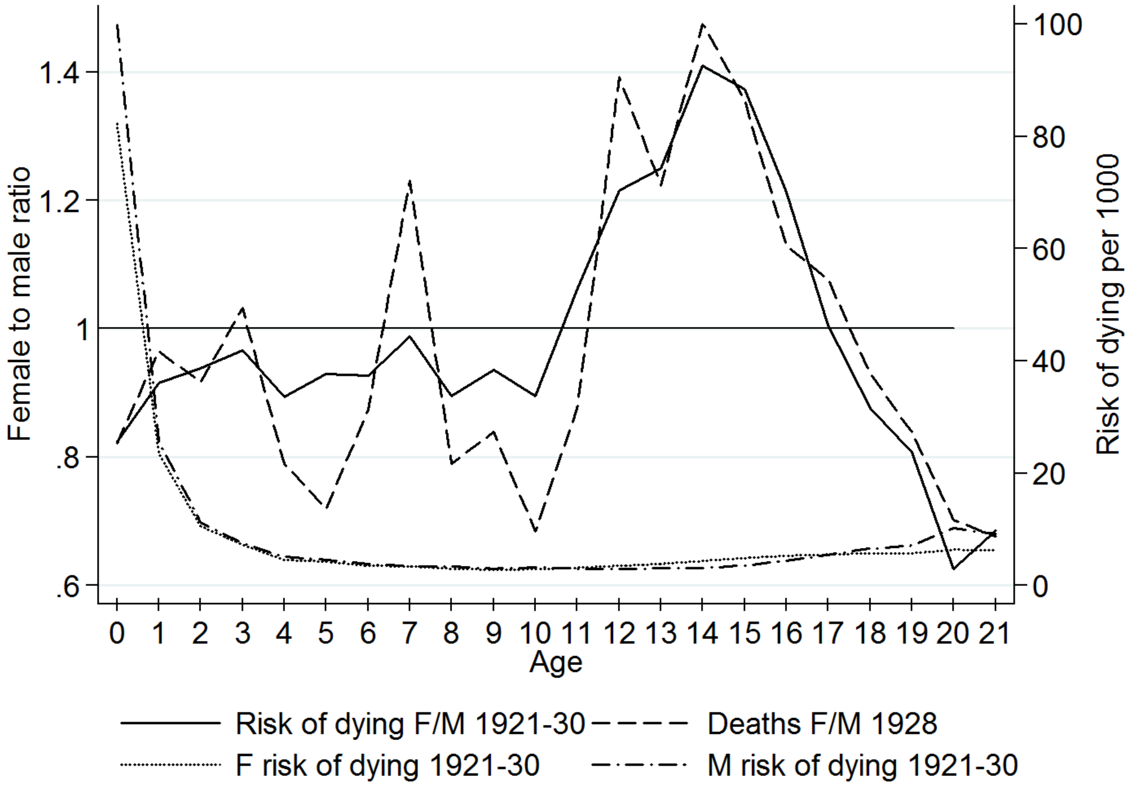

Figure 1. Indicators of sex differences in mortality. Female to male ratio for risks of dying in the 1920s and deaths in 1928 (left axis); and risk of dying per 1000 for each sex (right axis). Ratio of deaths in 1928 has been adjusted with a coefficient for the difference between ratios of death and ratios of mortality in 1930 (c. 1,03). *Sources:* Kannisto & Nieminen 1996, table 4A; SVT VI Väestötilastoa 75, table 19.

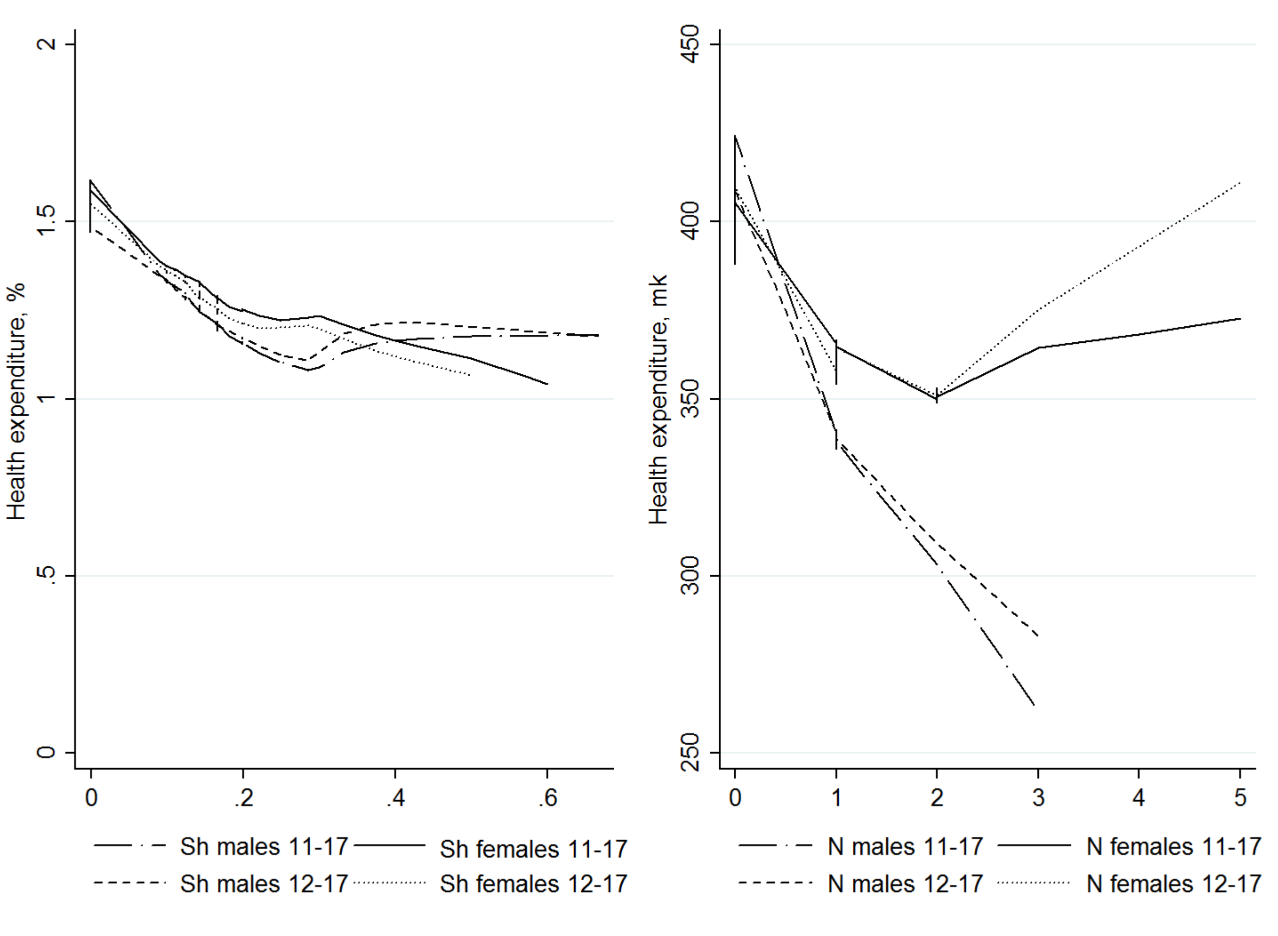

Figure 2. LOWESS lines on children by sex and health expenditure in the female excess mortality age groups in worker households. Shares (left) and amounts (right). Bandwidth 0.8. *Source:* CLS data.

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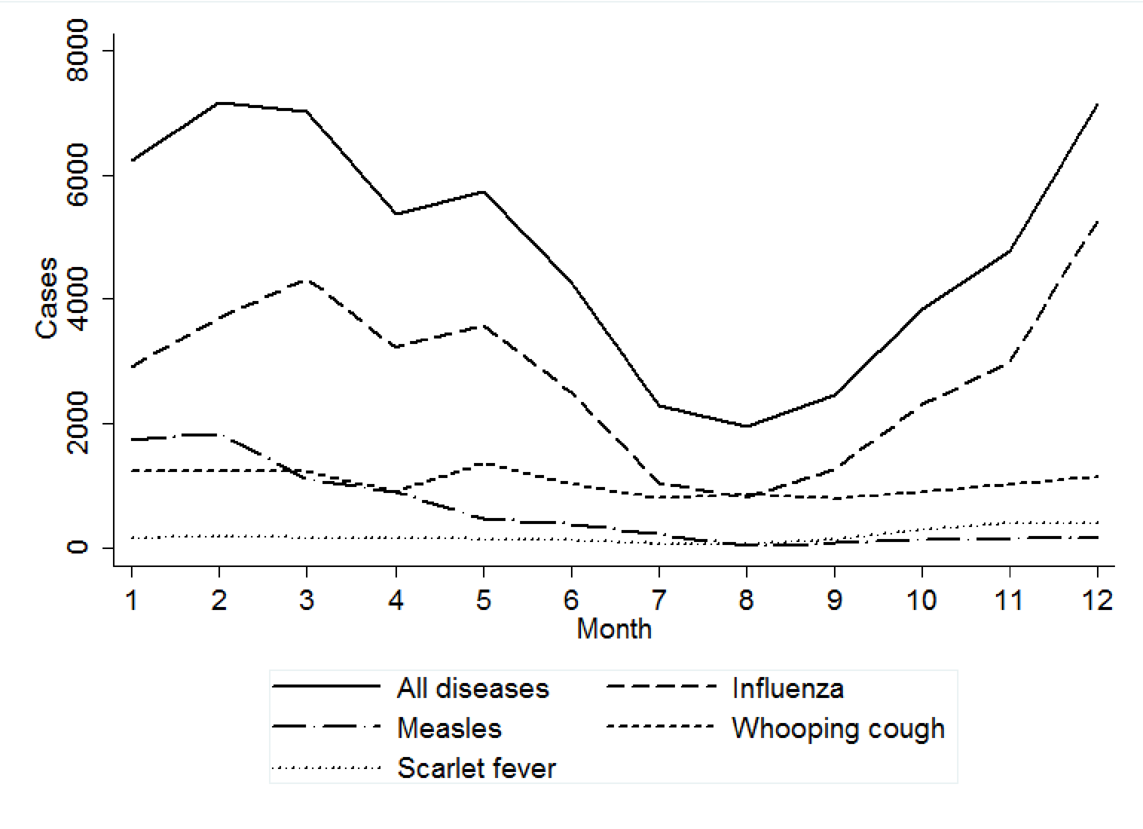


Figure 3. Monthly cases of contagious disease reported by district doctors in 1928.
Source: SVT XI Lääkintölaitos 45.

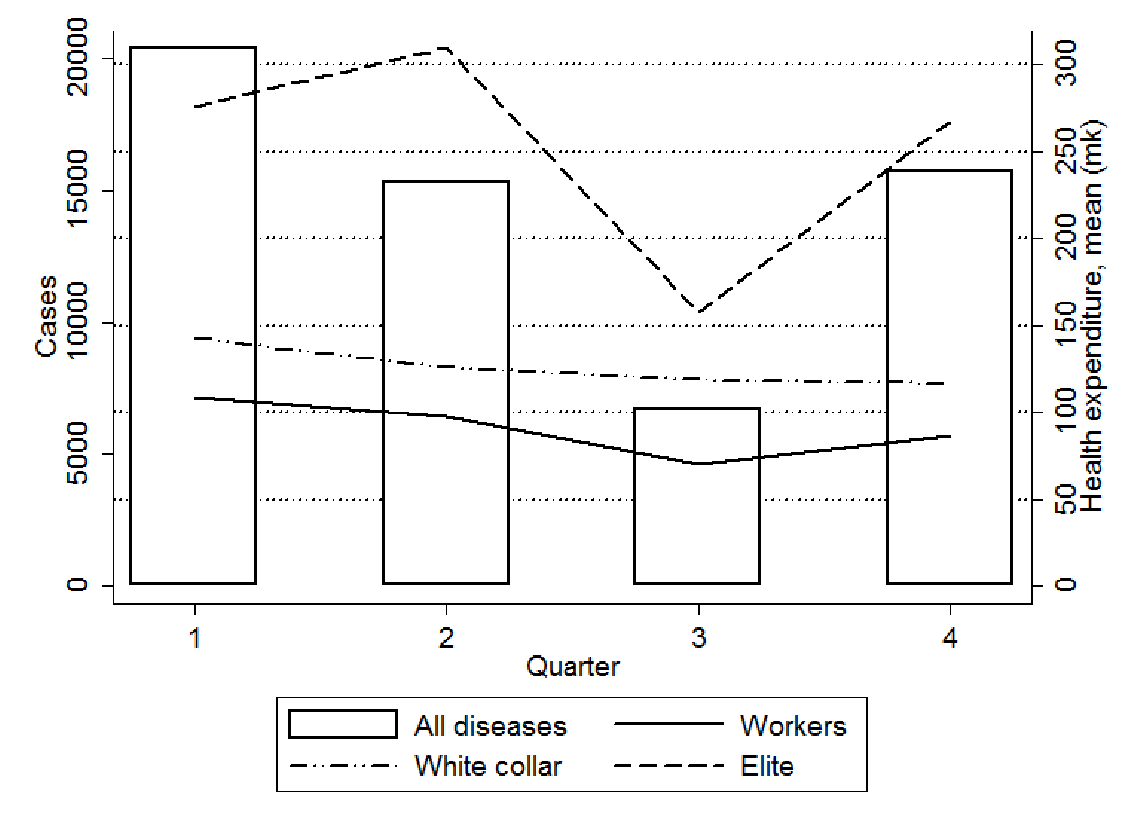


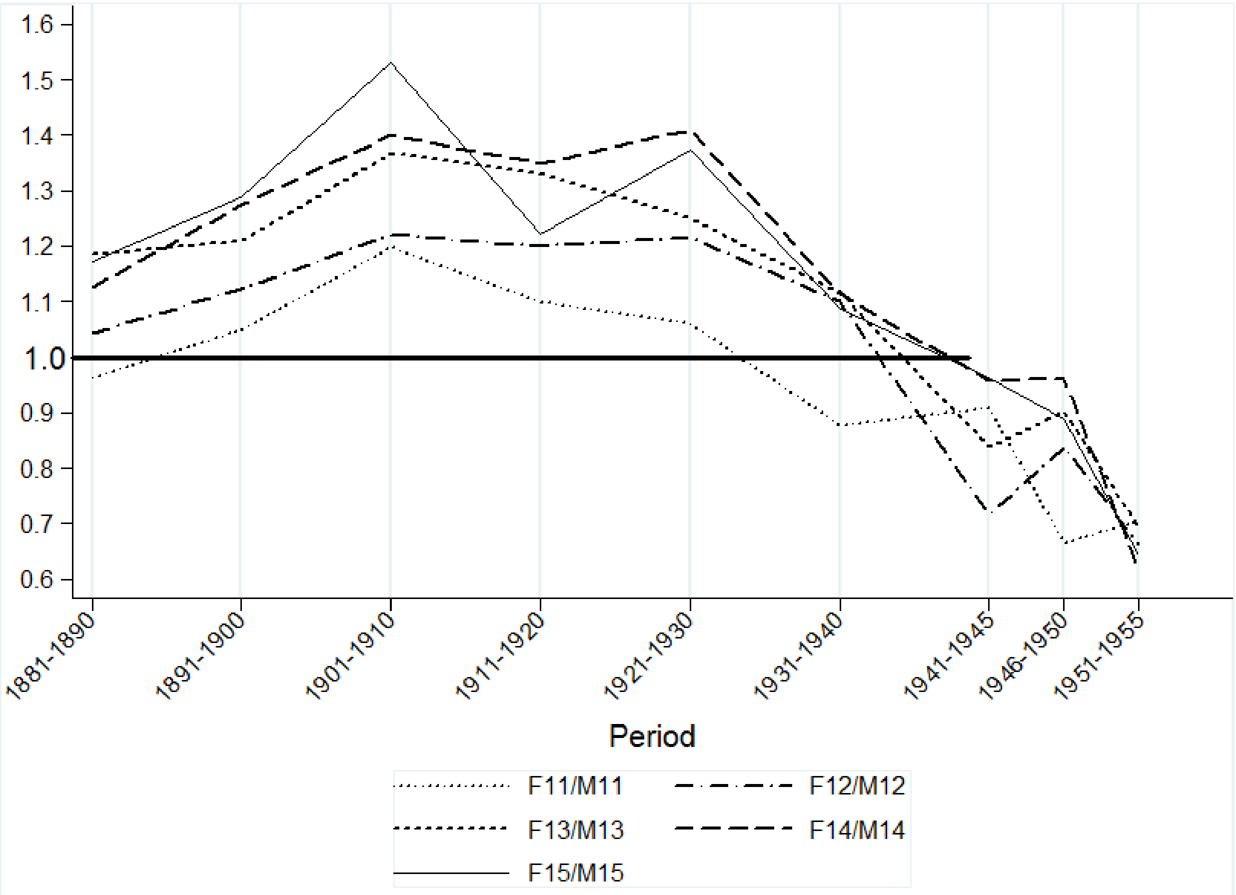
Figure 4. Cases of contagious disease reported by district doctors and the mean health expenditure of CLS households by quarter and by social group in 1928.
Sources: SVT XI Lääkintölaitos 45; CLS data.

Appendix tables and figures

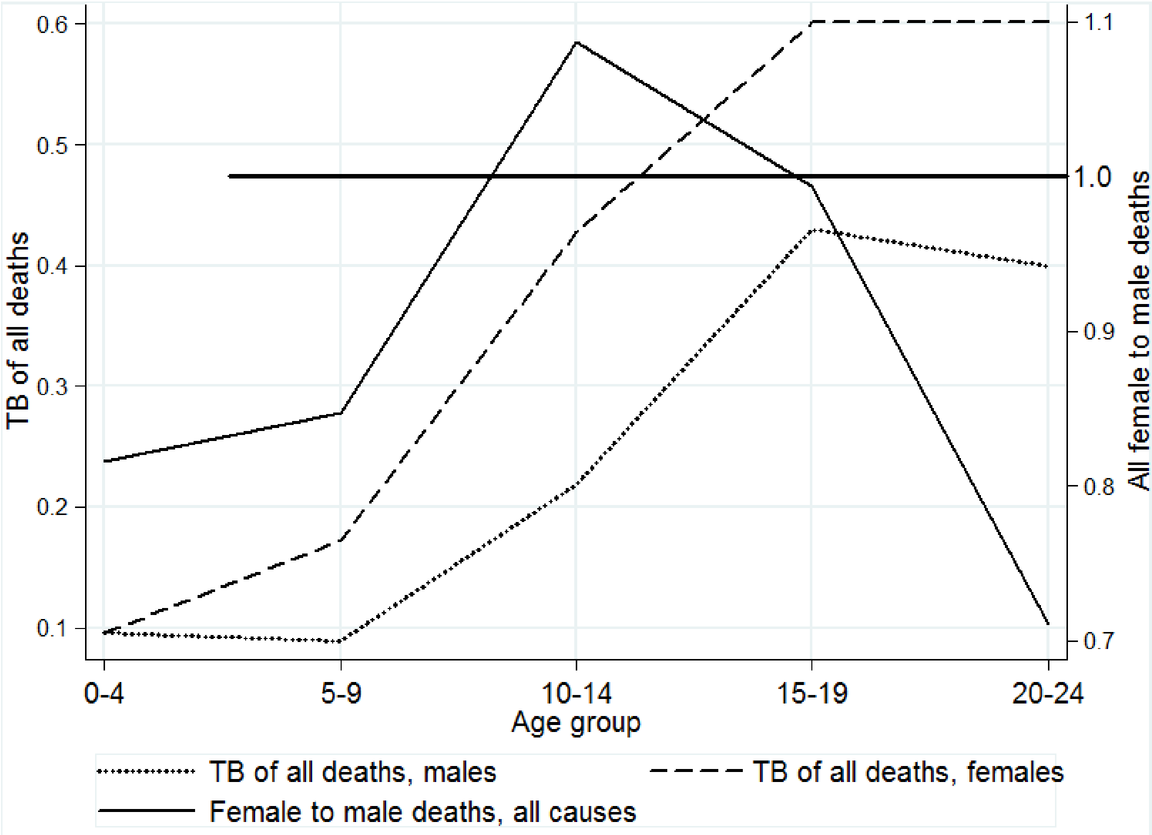
Model 1: Shares (Engel curves)				Model 2: Amounts			
Dependent: Health exp., % of total	Workers	White collar	Elite	Dependent: Health exp., mk	Workers	White collar	Elite
ln family size	-0.08 (0.63)	1.63 (1.66)	4.04 (3.42)	ln family size	552.34 (250.49)	894.68 (686.78)	4657.68 (2277.47)
ln exp per capita	0.25 (0.31)	1.15 (0.76)	1.64 (1.34)	ln exp per capita	260.55 (105.33)	498.33 (303.92)	1643.40 (780.76)
Share of females death risk 0-10	0.82 (0.74)	1.53 (1.68)	1.13 (1.39)	Females death risk 0-10	-0.82 (0.71)	1.02 (1.14)	3.44 (2.94)
Share of males death risk 0-10	1.29 (0.60)	2.03 (2.02)	1.11 (1.56)	Males death risk 0-10	-0.14 (0.41)	2.94 (1.72)	2.61 (2.45)
Share of females death risk 11-17	0.45 (0.88)	-0.65 (2.54)	-0.51 (2.91)	Females death risk 11-17	-5.26 (3.84)	-24.08 (7.92)	-27.00 (38.33)
Share of males death risk 11-17	-0.22 (0.91)	0.49 (2.97)	0.58 (5.32)	Males death risk 11-17	-16.20 (4.96)	-17.66 (14.08)	-30.40 (97.89)
Share of females death risk 18-29	1.31 (1.04)	0.39 (3.23)	-1.37 (3.80)	Females death risk 18-29	1.15 (4.42)	-2.88 (6.45)	-25.84 (34.06)
Share of males death risk 18-29	1.05 (0.82)	-0.64 (1.61)	1.91 (2.42)	Males death risk 18-29	1.35 (2.77)	-13.61 (5.19)	-13.80 (23.26)
Share of females death risk 30-54	1.69 (1.29)	3.96 (4.33)	3.44 (4.03)	Females death risk 30-54	-1.70 (6.43)	42.14 (17.04)	25.47 (26.66)
Share of females death risk over 55	1.91 (0.69)	1.39 (2.21)	1.11 (1.99)	Females death risk over 55	0.84 (0.42)	0.31 (1.10)	-6.38 (4.22)
Share of males death risk over 55	0.14 (0.77)	1.42 (1.49)	-0.72 (1.47)	Males death risk over 55	-3.64 (0.94)	0.20 (0.69)	-2.01 (8.13)
Constant	0.13 (1.52)	-5.04 (5.43)	-9.22 (9.28)	Constant	-1015.01 (558.89)	-2242.60 (1752.02)	-10104.34 (5281.81)
F-test gender diff 0-10	3.17*	1.08	0.00	F-test gender diff 0-10	1.78	2.32	0.07
F-test gender diff 11-17	1.39	0.94	0.03	F-test gender diff 11-17	4.82**	0.33	0.00
F-test gender diff over 55	14.01***	0.00	0.71	F-test gender diff over 55	16.28***	0.01	0.18
R^2	0.08	0.13	0.06	R^2	0.05	0.14	0.08
N	581	242	131	N	581	242	131

Appendix table 1. Health expenditure and risks of dying in age-sex category according to life tables 1921-1930. Shares of sum of total household members' risk of dying (left side) and risks of dying per 1000 (right side). OLS regressions by group.

Source: CLS data, computations from Kannisto & Nieminen 1996, table 4A. Standard errors clustered by locality in parentheses. ***=p<.01, **=p<.05, *=p<.10 (F-tests only).



Appendix figure 1. Female to male risk of dying by age from 11 to 15 years, 1880s-1950s. *Source:* Kannisto & Nieminen 1996, Table 4.



Appendix figure 2. Raw ratio of female to male deaths from all causes by age group and share of pulmonary tuberculosis of male and female mortality, 1928. *Source:* SVT VI Väestötilastoa 75, table 27.

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[1] E.g. for Helsinki the share of "workers" employed by the city itself was 15 %, although the entire workforce of the city would only have been 4 % of all those municipal taxpayers classifies as "workers" in published statistics. Helsingin kaupungin tilastokonttori 1929, 228-229; Helsingin kaupungin tilastotoimisto 1930, 252-255.

[2] The term, translated as "elite" in fact literally means "civil servant", but as this group included high-earner private sector employees as well, that translation would not be appropriate.

[3] The ratio for the 10-14 bracket available in the older life tables is roughly at parity in the 1870s, and shows higher male mortality from the 1820s through the 1860s (Turpeinen & Kannisto 1997, Table 1).

[4] If the limit between the younger and older children would be shifted from 11 to 12 years, as suggested by the 1928 data on deaths, all F-tests on children would fail. This is understandable for the younger group as no mortality data supports such grouping, and specifications focusing on narrower age brackets like 0-4 are robust (Saaritsa 2014, pp. 123-125). However, experiments with the older group show significant F-tests either for the Engel or amount specifications among workers only when using 11-17 or 11-15 as the bracket for female excess mortality, but not when using 11-16; and if using 12 as the lower boundary, only for the bracket 12-15.

[5] r=0.79 for males and 0.77 for females in the Engel specification, and about 0.54 with amounts.

[6] Regrettably, figures for many important conditions like measles are not available at a more disaggregated level.

[7] *Morbilli, tussis convulsiva, scarlatina, diphtheria, meningitis cerebrospinalis epidemica, poliomyelitis anterior acuta.*

[8] *Influenza, variola, typhus abdominalis, paratyphus, dysenteria, febris intermittens, encephalitis lethargica.*

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